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CHANGES IN THE RETURNS TO EDUCATION IN ARGENTINA

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In this paper we examine the returns to education in Argentina from 1995 to 2003. We use several econometric techniques in an attempt to account for sample selection bias arising from endogenous labour force participation and to control for the endogeneity of education. The empirical results indicate that the returns to education have fluctuated over time. We provide some evidence suggesting that the relative demand for more educated people is likely to be a key factor in explaining changes in the returns to education.

JEL classification codes: I21, J31
Key words: Argentina, returns to education, sample selection, endogeneity

I. Introduction

Argentina experienced dramatic structural reforms during the 1990s. The economy boomed over the 1990–1998 period, except for a brief downturn in 1995. Starting in 1998, however, Argentina entered into a period of recession that ended in the most severe economic crisis of its recent history in 2002. The whole period was characterized by increased volatility of output (Kehoe, 2003).

A number of researchers show that these economic fluctuations have had significant effects on labour market outcomes and the distribution of income. Galiani and Hopenhayn (2003) find that the number and the duration of unemployment spells have risen substantially in Argentina during the 1990s. This indicates an

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increase in unemployment risk and higher levels of job turnover in the labour market. This finding is in line with that obtained by Pessino and Andres (2003) who document the net job destruction experienced by Argentina between 1990 and 2001. As expected, job destruction was particularly pronounced between 1998 and 2001. During this period the proportion of individuals who lost their job and were unable to find a new one was, on average, about 19.6 per cent per semester. Finally, González Rozada and Menendez (2000) observe that the wage distribution widened between 1991 and 1998. Growing unemployment accounts for a large part of the increase in earnings inequality.

In line with these studies, we examine the evolution of the returns to education for the 1995-2003 period. Previous research has mainly analysed the relationship between earnings and education in Buenos Aires rather than in Argentina as a whole. Using 1985 data on Buenos Aires, Kugler and Psacharopoluos (1989) find that wage differentials by education were lower in Argentina relative to other Latin American countries. They suggest that the expansion of education is likely to have contributed to a relatively low education premium. However, since the mid-1980s the production process has undergone significant transformations that are likely to have had an impact on the demand for certain types of jobs and hence on the returns to education. This consideration is consistent with the results obtained by later studies. Pessino (1995) finds that the returns to education in Buenos Aires increased from 10 per cent in 1986 to 12.5 per cent in 1989, though they declined in 1990 but recovered again in 1993. A more recent study of the World Bank (2002) concludes that in Argentina well-educated workers enjoyed higher returns to education in 1998 than in 1992.

This paper also adds to the previous studies by using various econometric techniques to investigate the changes in the returns to education in Argentina. We begin by presenting standard OLS estimates of the Mincerian wage equation. Next we correct these estimates by accounting for the potential selectivity bias due to endogenous labour force participation. Finally, we deal with the endogeneity bias of education by employing an instrumental variable approach. Special attention is also given by this paper to the understanding of the main drivers behind the changes in the returns to education.

The remainder of the paper is organised as follows. Section II describes the data used for the estimation of the returns to education. In Section III we present and discuss the empirical findings. In Section IV we use a simple supply and demand framework in an attempt to explain the observed changes in the returns to education. Section V concludes.

II. Data

Ordinary least squares (OLS) methods are applied to simple human capital earnings functions of the Mincer-type:

$$\ln W_{it} = \alpha_{0t} + \alpha_{1t} ED_{it} + \alpha_{2t} EXP_{it} + \alpha_{3t} EXP_{it}^2 + \mu_{it}, \tag{1}$$

where ln W_{it} refers to hourly wage in logarithmic terms for individual i at time t, ED_{it} are years of schooling, EXP_{it} represents potential labour market experience, and μ_{it} is an independent and normally distributed error term with fixed variance σ_t^2 which does not need to be constant over time. Although results from this parsimonious specification are of a limited use, because of relatively low explanatory power and obvious omitted variable bias, they do compare nicely with results from other studies.

The data used in this paper come from a series of nine consecutive May waves (i.e. 1995-2003) of the Permanent Household Survey (EPH, Encuesta Permanente de Hogares - May).[1] This is a household labour force survey, undertaken by the National Institute of Statistics (INDEC), in provincial capitals and areas with more than 100,000 inhabitants.[2] It comprises questions on personal attributes, employment status, education, incomes as well as household characteristics.

The dependent variable of equation (1) is the net hourly wage resulting from the primary occupation of the individual. Following the definition given by the INDEC, hourly wages are constructed by dividing monthly earnings by actual working hours per week (multiplied by four). One should bear in mind that only net wages are reported and there is no information on the different tax rates levied on these wages. Although the survey records individual's education as a credential measure, following the approach of similar studies (see, for instance, Vieira 1999; Brunello and Miniaci 1999) it has been converted into a continuous indicator.[3] More specifically, years of education are calculated according to the following procedure.

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[1] Unfortunately we cannot examine changes in the returns to education over a longer period of time, as data on the early 1990s are unavailable.

[2] We use data for all those urban conglomerates for which information is available in each survey year. Although one should be aware that more recent waves of the EPH include new regions and cities, we do not believe this introduces a bias in our results given that the large majority of data continue to come from those urban conglomerates that are common to all survey years.

[3] In constructing a continuous measure of years of education it is assumed that returns to education are linear. In an attempt to provide some evidence on the extent of the bias introduced by imposing such a restriction on the shape of the earnings-schooling profile, following Patrinos (1996) we estimate a wage equation allowing for a more flexible education function. Thus education is represented by dummy

First, we compute the nominal number of completed years in order to achieve the highest educational level reported by the individual.[4] Second, as the survey gives information on the number of years spent by the individual at a given educational level, though he/she did not complete it, we sum this number to the first one. As the EPH does not provide any information on either actual work experience or years of work interruption, we use potential work experience (defined as age - years of schooling - 6) as an explanatory variable in the earnings function. Note that potential work experience is also included as a quadratic term in the wage specification in order to capture the concavity of the experience-earnings profile.

Our analysis focuses on full-time employees aged between 14 and 65.[5] Following the approach of Hægeland et al. (1999) a person is defined a full-time worker if he/she works more than 30 hours per week. In line with similar studies (see, for instance, Edin and Holmund 1993 and Fersterer and Winter-Ebmer 2003) self-employed workers and those without pay are excluded from the final sample.

III. Results

A. Standard OLS estimates

Table 1 presents standard OLS estimates of the Mincerian earnings equation for the combined sample of men and women. Not only do we include a gender dummy variable, but we also interact this with education, potential work experience and its square. Returns to education have fallen between 1996 and 1998. Since then, however, the trend reversed as average returns to one additional year of education increased by approximately one-tenth, from 8.48 per cent in 1998 to 9.26 per cent in 2002. Finally, the year 2003 saw a decline in the rate of return to education with this measure decreasing to 8.64 per cent.[6]

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variables for each year of education (the education variable begins at 8 years and it is truncated at 19 years). Although these estimates (not reported due to space reasons) show that returns to education tend to increase discontinuously, we still choose to employ a continuous measure of education. This is in order to make our results comparable to the large majority of studies that use this approach.

[4] The statutory number of years required to obtain a primary school degree and a secondary school degree is 7 and 5 respectively. Successfully completing tertiary education typically requires 5 years.

[5] The Law on Labor Contracts sets the minimum age for employment at 14 years.

[6] The differences in returns to education between 1996 and 1998, 1998 and 2002, and 2002 and 2003 are all statistically significant.

Table 1. Estimates of basic Mincerian equation, male and female employees

Year	1995	1996	1997	1998	1999	2000	2001	2002	2003
Const	-0.4365***	-0.5755***	-0.4912***	-0.4760***	-0.4849***	-0.5371***	-0.5920***	-0.7104***	-0.5712***
	(0.0222)	(0.021)	(0.0209)	(0.021)	(0.0233)	(0.0258)	(0.0273)	(0.0318)	(0.0342)
FEMALE	-0.1024***	-0.0790**	-0.0588	0.0307	-0.1422***	-0.0278	0.0441	0.0435	-0.0293
	(0.0394)	(0.0373)	(0.0369)	(0.0367)	(0.0402)	(0.0438)	(0.0461)	(0.0529)	(0.0591)
EXP	0.0488***	0.0497***	0.0504***	0.0534***	0.0511***	0.0470***	0.0486***	0.0513***	0.0479***
	(0.0015)	(0.0014)	(0.0012)	(0.0015)	(0.0018)	(0.0014)	(0.0019)	(0.0021)	(0.0023)
EXP*FEMALE	-0.0059***	-0.0066***	-0.0085***	-0.0110***	-0.0041	-0.0100***	-0.0077***	-0.0106***	-0.0042
	(0.0025)	(0.0024)	(0.0024)	(0.0025)	(0.0027)	(0.0029)	(0.003)	(0.0034)	(0.0037)
EXP^2	-0.0007***	-0.0006***	-0.0007***	-0.0007***	-0.0006***	-0.0007***	-0.0006***	-0.0006***	-0.0005***
	(0.00003)	(0.00003)	(0.00003)	(0.00003)	(0.00003)	(0.00004)	(0.00004)	(0.00004)	(0.00004)
EXP^2*FEMALE	0.00004	0.0001	0.0001	0.0001	0.00001	0.00001	0.00005	0.00001	0.00002
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
ED	0.0921***	0.0971***	0.0887***	0.0848***	0.0853***	0.0864***	0.0886***	0.0926***	0.0864***
	(0.0014)	(0.0013)	(0.0012)	(0.0012)	(0.0013)	(0.0014)	(0.0015)	(0.0017)	(0.0019)
ED*FEMALE	0.0021	0.0013	0.0020	-0.0017	0.0054***	-0.0001	-0.0038	-0.0019	-0.0015
	(0.0024)	(0.0023)	(0.0021)	(0.002)	(0.0023)	(0.0024)	(0.0025)	(0.0028)	(0.0032)
Num.obs.	17,850	20,613	20,853	20,421	17,422	15,162	14,930	12,627	9,787
Adj. R^2	0.3193	0.3297	0.3121	0.3039	0.3063	0.3010	0.2833	0.2904	0.2811

Notes: standard error in parentheses; * Significant at 10 percent level; ** Significant at 5 percent level; *** Significant at 1 percent level.

These estimates are not fully consistent with those obtained by Giovagnoli et al. (2004) who find that the returns to education in Argentina exhibited a positive trend throughout the period under examination. This conflicting outcome could be attributed to substantial differences in the way the final sample was constructed. Unlike this study, Giovagnoli et al. (2004) do not exclude self-employed workers from their analysis and do not focus only on full-time workers. Thus it is possible that during the economic expansion of 1996 to 1998 the higher returns to education for self-employed and part-time workers could have driven the overall returns to schooling up. Additionally, the analysis carried out by Giovagnoli et al. (2004) covers a smaller number of urban conglomerates relative to our study.[7]

Note also that the explanatory power of the wage equations, measured in term of adjusted R^2, tends to fall after 1997. Therefore, omitted variables are likely to play a greater role in explaining changes in earnings in more recent years. This may simply reflect increased (observed and unobserved) workers' heterogeneity.

For international comparisons, we can observe that returns to education are higher in Argentina than in most industrialised countries (Trostel et. al. 2002). This finding supports the hypothesis that the rate of return to education is negatively correlated with per capita income.

B. Sample selection

The estimates shown in Table 1 are likely to suffer from sample selection bias. The sample used in the estimation of the wage equation comprises only full-time workers. Selection into full-time employment may not be random as full-time workers may have systematically different characteristics from those without a full-time job. If

[7] We have also calculated wage estimates for men and women separately. Although this is a standard approach in many studies on returns to education, one should be cautious about drawing conclusions from these estimates because results for males and females may not be fully comparable. Labor turnover is typically higher among female workers relative to male workers since women's decision to participate in the labour market is often significantly affected by non-economic reasons (for a discussion on this issue in Argentina, see, for instance, Paz 2004). OLS estimates indicate that in the first half of the period under examination the returns to education were higher for women than for men, while the opposite is true for the second half of the period. It is possible that in the recession of 1999-2002 the substantial increase in female labour participation has been accompanied by an increase in the relative supply of more educated women, leading to a decline in the rate of return to education. Higher unemployment amongst male heads of households during the economic crisis has encouraged a large number of educated women to seek employment. Finally, in the remainder of the paper we only report and discuss estimates for the combined sample of men and women.

these characteristics exert an influence not only on the individuals' probability of having a full-time job but also on their earnings potential, then a full-time employment selection equation needs to be considered when estimating the wage equation.

There is also an additional reason why in our analysis it might be especially relevant to estimate a model that includes an employment equation. As outlined in the introduction, between 1995 and 2003 Argentina experienced a significant increase in the unemployment rate as well as in the number and duration of unemployment spells. These factors might have had an impact on returns to education by affecting, for instance, working hours (Ashenfelter and Ham 1979).

One way of accounting for this potential bias is to use the standard 'two-step' estimation method proposed by Heckman (1979). The first step consists in estimating a probit model for full-time labour force participation. Let y^*_{it} be a latent variable indicating the propensity of the individual i to have a full-time job at time t. This can be expressed as:

$$y^*_{it} = \beta_{ot} X_{it} + \varepsilon_{it}, \tag{2}$$

where X_{it} is a vector containing individual characteristics that are thought to influence full-time labour force participation and ε_{it} is a random error term.

Note that y^*_{it} is unobserved. Instead we observe a dichotomous variable, y_{it}, that is equal to 1 if $y^*_{it} > 0$ while it takes the value of 0 if $y^*_{it} \leq 0$.

From the estimation of equation (2) we compute the inverse Mill's ratio that is included as an additional regressor in the wage equation. Thus equation (1) becomes:

$$\ln W_{it} = \alpha_{0t} + \alpha_{1t} ED_{it} + \alpha_{2t} EXP_{it} + \alpha_{3t} EXP^2_{it} + \alpha_{4t} \lambda_{it} + \mu_{it}, \tag{3}$$

where $\lambda_{it} = \phi(\beta_{ot} X_{it}) / \Phi(\beta_{ot} X_{it})$, $\phi(.)$ is the density standard normal distribution and $\Phi(.)$ is the cumulative standard distribution.

Improving parameter specification requires at least one valid instrument, i.e., a variable affecting full-time labour force participation, but that does not have any impact on wages. Following the approach of Duraisamy (2000), our instrument is a dummy variable recording whether the individual has any source of non-labour income. The rationale for this is that one would expect non-labour income sources to discourage people from having a full-time job. Additionally, given the greater family commitment usually borne by women relative to men within households, we also use the presence of small children to identify female labour force participation. Thus, in line with several studies (see, for instance, Betts 2001), we add to the full-time employment selection equation a dummy variable recording whether the woman has at least one small child (aged less or equal to 2).[8,9]

Table 2 shows the results of the earnings equations corrected for selection into full-time employment. The coefficients on the selection term are statistically significant at all conventional levels. Such result supports the appropriateness of the Heckman's technique and hence suggests that the estimates obtained from a standard OLS method are likely to be inefficient. Although the negative coefficient of λ indicates a negative selection into full-time employment, this is not an unusual finding in the labour economics literature (see, for instance, Nicaise 2001). One possible explanation is that an increase in the returns to education has encouraged many highly educated full-time workers to reduce the number of working hours. This interpretation is also consistent with the fact that the sub-period 1995-1996 is characterised by both the highest rates of return to education and the highest values of the coefficient of the selection term. Although standard OLS estimates of the rate of return to education appear to be slightly upward biased relative to those controlling for selection into full-time employment, they show a similar inter-temporal pattern.[10]

C. Endogeneity of education

Research in education economics suggests that standard OLS estimates of the return to schooling may be biased because of endogeneity of education. Endogeneity can arise as result of an individual's optimal schooling choices, measurement error and omitted variables (Harmon et al. 2003).

The instrumental variable (IV) approach has often been used in the economic literature to address the endogeneity bias. In order to employ this method one needs at least one valid instrument, i.e., a variable that is correlated with schooling, but that has no impact on wages. The IV estimation technique comprises two stages.

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8 Unfortunately all the identifying variables for the full-time employment selection equation are dummies. This may pose a problem given that dummy variables might not have enough variation to constitute a good instrument (Heckman 1990).

9 Thus in the final regression our identifying instruments are: an interaction term between a female dummy variable and the presence of at least one small child, and the non-labour income dummy variable.

10 Differences in the rate of return to education between standard OLS estimates and OLS estimates corrected for selection into full-time employment are found to be statistically significant at all conventional levels (see Appendix for details).

Table 2. Estimates of basic Mincerian equation corrected for sample selection bias, male and female employees

Year	1995	1996	1997	1998	1999	2000	2001	2002	2003
Const	-0.1923***	-0.3189***	-0.3976***	-0.3983***	-0.3958***	-0.4495***	-0.4494***	-0.5089***	-0.4013***
	(0.0412)	(0.0408)	(0.0351)	(0.0343)	(0.0385)	(0.0434)	(0.0477)	(0.0591)	(0.0627)
FEMALE	0.0355	0.0528	-0.0054	0.0722*	-0.0917**	0.0195	0.1154**	0.1216**	0.0456
	(0.0439)	(0.0413)	(0.0402)	(0.0395)	(0.0438)	(0.0477)	(0.05)	(0.0563)	(0.06343)
EXP	0.0379***	0.0388***	0.0466***	0.0503***	0.0458***	0.0478***	0.0434***	0.0449***	0.0422***
	(0.0021)	(0.002)	(0.0018)	(0.0018)	(0.002)	(0.0022)	(0.0023)	(0.0027)	(0.0029)
EXP*FEMALE	-0.0068***	-0.0071***	-0.0087***	-0.0110***	-0.0043***	-0.0104***	-0.0080***	-0.0110***	-0.0046***
	(0.0025)	(0.0024)	(0.0024)	(0.0025)	(0.0027)	(0.0029)	(0.003)	(0.0034)	(0.0037)
EXP^2	-0.0004***	-0.0004***	-0.0006***	-0.0007***	-0.0006***	-0.0006***	-0.0005***	-0.0005***	-0.0005***
	(0.00004)	(0.00004)	(0.00004)	(0.00004)	(0.00004)	(0.00004)	(0.00005)	(0.0001)	(0.0001)
EXP^2*FEMALE	0.0001	0.0001	0.0001	0.0001	0.00002	0.0002**	0.0001	0.0001	0.00003
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
ED	0.0878***	0.0922***	0.0868***	0.0832***	0.0833***	0.0845***	0.0854***	0.0879***	0.0826***
	(0.0015)	(0.0015)	(0.0014)	(0.0013)	(0.0015)	(0.0016)	(0.0018)	(0.0021)	(0.0022)
ED*FEMALE	0.0004	-0.0005	0.0011	-0.0023	0.0047**	-0.0008	-0.0049	-0.0028	-0.0025
	(0.0024)	(0.0023)	(0.0021)	(0.0021)	(0.0023)	(0.0024)	(0.0025)	(0.0028)	(0.0032)
λ	-0.1699***	-0.1668***	-0.0631***	-0.0546***	-0.0608***	-0.0571***	-0.0909***	-0.1117***	-0.0971***
	(0.0242)	(0.0227)	(0.019)	(0.019)	(0.0209)	(0.0227)	(0.0249)	(0.0276)	(0.03)
Num. obs.	17,850	20,613	20,853	20,421	17,422	15,162	14,930	12,627	9,787
Adj. R^2	0.3211	0.3314	0.3125	0.3042	0.3066	0.3012	0.2839	0.2912	0.2818

Notes: standard errors in parentheses; * Significant at 10 percent level; ** Significant at 5 percent level; *** Significant at 1 percent level.

First, we estimate the instrumental variable against schooling and we retain the predicted values for schooling. In the second stage we estimate the earnings function where we replace actual schooling with the predicted values for schooling obtained in the first stage.

Following the approach of Trostel et al. (2002), we use spouse's education as our instrument. The intuition behind this variable is that years of education of husbands and those of wives are positively correlated. Individuals with comparable levels of education tend to get together as they are likely to share similar experiences and/or similar interests. This hypothesis is supported by a number of empirical works (see, for instance, Pencavel 1998). On the other hand, it is not obvious that spouse's education exerts a direct influence on individual's earnings.

Our choice of the instrumental variable has been determined by data availability. It is very hard to find another instrument for education within the EPH dataset. For instance, since the EPH does not provide any information on parental characteristics, father's education and mother's education cannot be used for this purpose. One may also note that the use of spouse's education is accompanied by two main drawbacks. First, IV estimates are based on a considerably smaller sample size relative to those shown in Tables 1 and 2. Obviously we can only include in the final sample those individuals for whom data on spouse's education are available. Second, selection into marriage may not be random. It is possible that some of those unobservables affecting the choice of getting married may also have an impact on individual's earnings, thereby leading to biased estimates. Although these limitations may jeopardize the comparability of the IV estimates with the previous ones, we still feel that it is important to attempt to control for the endogeneity of education and thus the IV results are reported.

So far in this subsection we have only discussed the issue of the endogeneity of education. However, if the number of years spent in school is endogenous, then also potential work experience, which has been defined as age – years of schooling -6, is endogenous. In an attempt to address this problem, in the wage equation we replace potential work experience and its squared with age and age squared, respectively. This is because age can be safely treated as an exogenous variable (Uusitalo 1999).

Table 3 depicts the IV estimates of the Mincerian earnings equation where we account for the endogeneity of both education and potential work experience. The IV estimates of the rate of return to education are found to be considerably higher than the corresponding OLS estimates throughout the whole period. This downward bias in the OLS estimates is in line with the results obtained by many previous

Table 3. IV estimates of basic Mincerian equation, male and female employees

Year	1995	1996	1997	1998	1999	2000	2001	2002	2003
Const	-1.0232***	-1.3034***	-1.0489***	-1.0726***	-1.0028***	-1.5366***	-1.1346***	-1.3123***	-1.2105***
	(0.1435)	(0.1365)	(0.1467)	(0.1904)	(0.1523)	(0.1535)	(0.1556)	(0.1751)	(0.2038)
FEMALE	0.2924	-0.3314	-0.0963	0.0131	-0.4199	0.3680	0.1525	-0.0350	-0.3999
	(0.3049)	(0.2954)	(0.3219)	(0.3766)	(0.3207)	(0.3118)	(0.3215)	(0.353)	(0.427)
AGE	0.0533***	0.0534***	0.0379***	0.0357***	0.0417***	0.0594***	0.0411***	0.0415***	0.0323***
	(0.0071)	(0.0066)	(0.0073)	(0.0097)	(0.0077)	(0.0077)	(0.008)	(0.0087)	(0.0102)
AGE*FEMALE	-0.0184	0.0214	0.0211	0.0251	0.0107	-0.0211	-0.0193	-0.0005	0.0182
	(0.0151)	(0.0146)	(0.0163)	(0.0187)	(0.0165)	(0.0156)	(0.0165)	(0.0172)	(0.021)
AGE^2	-0.0005***	-0.0005***	-0.0003***	-0.0002***	-0.0003***	-0.0005***	-0.0003***	-0.0003***	-0.0002***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
AGE^2*FEMALE	0.0001	-0.0004**	-0.0004**	-0.0005***	-0.0002	0.0002	0.0002	-0.0001	-0.0003
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0003)
ED	0.1007***	0.1159***	0.1187***	0.1312***	0.1043***	0.1189***	0.1112***	0.1149***	0.1259***
	(0.0051)	(0.005)	(0.0056)	(0.0087)	(0.0061)	(0.0059)	(0.0063)	(0.0067)	(0.0089)
ED*FEMALE	0.0008	-0.0071	-0.0273***	-0.0473***	0.0165	-0.0043	0.0128	0.0055	0.0013
	(0.0095)	(0.0094)	(0.0104)	(0.0139)	(0.0117)	(0.0112)	(0.0117)	(0.0124)	(0.0166)
Num.Obs	5,839	6,056	5,413	3,290	4,909	4,968	5,101	4,549	3,380
Adj R^2	0.1109	0.1374	0.1284	0.1225	0.1209	0.1411	0.1234	0.1221	0.1206

Notes: standard errors in parentheses; * Significant at 10 percent level; ** Significant at 5 percent level; *** Significant at 1 percent level.

studies (Brunello and Miniaci 1999; Harmon and Walker 1995). In contrast to the OLS estimates, IV estimates indicate that the rate of return to education increased between 1995 and 2003.

The empirical validity of the IV estimates rests on the instrument being a significant determinant of schooling but not of earnings. In line with Bound et al. (1995), we run two different tests to check the 'quality' of our instrument. In the first half of Table 4 we report the t-values associated with the coefficient on the spouse's education in the schooling equation. It can be seen that the instrument is a good predictor of schooling for all the years under examination. Second, we regress the wage residuals from the IV estimation against the instrument in order to ensure that the instrument is not directly correlated with the wage once we control for other explanatory factors. The t-values associated with the coefficient on the instrument are depicted in the second half of Table 4. Our results indicate that throughout all the years the instrument does not explain any significant variation in the residuals. Thus both tests support the hypothesis that our instrument is 'valid'.

In an attempt to bring the OLS and IV estimates together, the following consideration can be made. Although the IV estimates exhibit a different inter-temporal pattern relative to the OLS estimates, these econometric methods concur that the rate of return to education was higher in 1996 than in 1999. Similarly, the rate of return to education increased in the 1999-2002 period.

In the next Section we attempt to shed light on the possible reasons for these changes in the education premium. This is a difficult exercise given that, although there is

Table 4. Instrument validity test, male and female employees

	(1)	(2)
1995	39.21***	0.95
1996	40.03***	0.12
1997	34.37***	0.51
1998	22.42***	0.39
1999	30.63***	0.59
2000	31.28***	0.55
2001	30.00***	-0.13
2002	29.08***	0.73
2003	21.31***	0.76

Notes: (1) is the t-value associated with the instrument in the schooling equation; (2) is the t-value associated with the instrument when wage residuals from the IV estimation are regressed against the instrument; *** Significant at 1 percent level.

evidence that the rate of return to education was higher in both 1996 and 2002 relative to 1999, the size of these differences is relatively small. Nevertheless, little variation in the returns to education seems to be a common problem in several studies (see, for instance, Vieira 1999) confined to the analysis of a relatively short period of time.

IV. Changes in the returns to education: Supply and demand

It is important to investigate the extent to which changes in the returns to education can be explained using a simple supply and demand framework. Let's first focus on shifts in the relative supply of more educated people. One simple way of doing this is to compare changes in the rate of return to education with changes in the average years of education for individuals aged 14 to 65 during the periods 1996-1999 and 1999-2002. As shown in Table 5, these variables move in the same direction. They decreased between 1996 and 1999, while they increased between 1999 and 2002. The positive relationship between the relative supply and the returns to education appear not to be consistent with a simple model with stable relative demand and fluctuating relative supply. If the relative demand for more educated people is stable over time, an increase in the relative supply of more educated people is in fact expected to lower the returns to education while a decline in the relative supply is supposed to bring them up.

As shifts in the relative supply seem not be the driving force behind the observed changes in the returns to education, we turn our attention to changes in the relative demand for labour. A significant number of studies (see, for instance, Katz and Murphy 1992) show that changes in the relative demand for more educated workers play a pivotal role in explaining the rise in wage inequality experienced in the United States throughout the 1980s. One reason for improved wage prospects of relatively skilled workers lies in the impact of new technology – the so called 'skill-biased technological' hypothesis (Machin and Van Reenen 1998).

Table 5. Returns to education and relative supply, male and female employees

Period	Changes in average years of education for individuals aged 14 to 65 (%)	Changes in the rate of return to education (%)		
		Standard OLS estimates	OLS estimates corrected for sample selection bias	IV estimates
1996-1999	-6.41	-12.22	-9.71	-10.00
1999-2002	3.16	8.59	5.50	10.15

Table 6 depicts changes in the full-time employment distribution across fifteen industries and three occupational groups in Argentina in the 1995-2003 period. These shifts may provide an indication of the growth in the relative demand for labour favouring more-educated workers over the less-educated ones. To better illustrate this phenomenon, we group the industries into three categories according to the expected skill level of most their workers (i.e., high skill, medium skill and low skill). Next we calculate the rate of growth of the share of employment accounted by each of these categories over between 1996 and 1999 as well as between 1999 and 2002. These values are reported at the bottom of Table 6.

The evidence presented in Table 6 suggests that the demand for less-educated workers is likely to have exhibited a positive change between 1996 and 1999 as a result of an increase in the share of employment accounted by industries that intensively employ less-skilled workers. By contrast, in the same period there seems to be evidence of a very small decline in the demand for more-educated workers. Relative employment in high skill intensive sectors, such as education and banking, fell slightly. Changes in the occupational distribution of employment also support the hypothesis that the 1996-1999 period has been characterised by a little drop in the demand for more-educated workers. This is shown by a decline in importance of professionals and managers. Therefore, the combination of lower demand for more-educated workers and especially of higher demand for less-educated workers could have depressed the returns to education.

The opposite situation is likely to have occurred between 1999 and 2002. In this period shifts in the employment distribution from low skill intensive sectors to high skill intensive sectors seem to indicate a demand shift trend in favour of more-educated workers and against the less-educated ones. This conclusion is also corroborated by the shifts in the occupational distribution of employment. The rising demand for more-educated workers and the falling demand for less-educated workers could have yielded an increase in the returns to education.

To more formally analyse the role of relative supply and demand changes in explaining relative wage changes, we adopt the approach of Katz and Murphy (1992), which has also been used in other studies (see, for instance, Fersterer and Winter-Ebmer 2003 and Giovagnoli et al. 2004). Thus we divide our data into 60 distinct labour groups, distinguished by sex, six education categories (0-5, 6-8, 9-12, 13-15, 16-17, 18+ years of schooling) and five experience categories (0-8, 9-16, 17-24, 25-32, 33-40 years of experience). For these groups, we calculate the changes in supply (Δs) and the changes in mean wages (Δw). In Figures 1 and 2 we plot the changes in relative wages against the changes in relative supply during

Table 6. Average industrial and occupational employment distribution, male and female employees

Percentage employment shares	1995	1996	1997	1998	1999	2000	2001	2002	2003
Agriculture	2.68	2.63	2.58	2.19	1.98	2.21	2.41	2.52	3.24
Textile, clothing, tobacco and food	6.22	6.83	6.49	6.31	5.51	5.33	5.21	5.59	6.12
Chemical manufactiring	1.52	1.66	1.55	1.66	1.70	1.63	1.46	1.65	1.25
Machinery and electrical equipment	4.11	3.50	3.73	3.47	3.41	3.21	2.65	2.23	2.39
Other manufacturing	3.24	3.34	3.41	3.20	3.00	3.08	2.92	2.63	3.09
Communication, transport and utilities	9.17	9.04	8.83	9.11	9.21	9.36	9.54	9.75	9.16
Construction	6.87	6.24	7.46	8.88	9.03	7.01	7.00	4.65	5.72
Wholesale trade	4.30	4.37	4.36	4.24	4.71	4.92	5.10	4.14	4.25
Retail trade	9.27	9.68	9.76	10.38	10.73	10.82	10.92	11.23	10.87
Hotels and restaurants	2.91	2.65	2.49	2.63	2.55	2.71	3.25	2.51	2.41
Banking	2.82	2.51	2.69	2.41	2.53	2.67	2.57	2.82	2.46
Medical and business services	10.18	10.59	10.33	10.47	10.81	11.24	11.26	11.96	11.42
Public adminstration and defense	19.57	19.96	19.63	18.10	17.43	17.78	17.91	19.45	19.35
Public education	5.27	5.23	5.05	5.07	4.92	5.24	5.30	6.31	6.29

Table 6. (continued) Average industrial and occupational employment distribution, male and female employees

Percentage employment shares	**1995**	**1996**	**1997**	**1998**	**1999**	**2000**	**2001**	**2002**	**2003**
Other services	11.89	11.78	11.66	11.89	12.48	12.78	12.51	12.56	11.98
Professionals, technical and managers	9.67	9.34	9.28	8.39	8.33	8.97	8.86	9.21	8.40
Sales and clerical	13.95	14.18	18.22	19.59	19.39	20.35	20.15	21.96	22.40
Production and service workers	76.37	76.49	72.51	72.02	72.28	70.68	70.99	68.83	69.19
Rate of growth of employment shares (percentage)		**1996-1999**			**1999-2002**				
Low skill intensive industries*		2.53			-21.14				
Medium skill intensive industries**		-0.65			2.09				
High skill intensive industries***		-0.39			15.47				

Notes: * comprises: agriculture, textile, clothing, tobacco, food, construction, other manufaturing; ** comprises chemical manufacturing, machinery and electrical equipment, wholesale trade, hotel and restaurants, public adminstration and defence, communication, transport and utilities, retail trade, and other services: *** comprises banking, medical and business service and public education.

the periods 1996-1999 and 1999-2002, respectively. If the demand for skills remains constant over time, one would expect the evolution of wages to be negatively related to changes in supply. This, however, does not occur in Figures 1 and 2 as they both show a significantly positive relationship between changes in supply and changes in wages. The shown regression lines are (standard errors in parentheses):

- For the 1996-1999 period: $\Delta w = -0.046 + 0.353\ \Delta s$ (0.014) (0.044) $R^2 = 0.528$

- For the 1999-2002 period: $\Delta w = -0.083 + 0.182\ \Delta s$ (0.017) (0.051) $R^2 = 0.181$

Although this is a simple framework in which one cannot isolate the exact contribution of supply and demand changes, it emerges that movements in the demand for skills are likely to be a key factor in explaining the observed changes in the returns to education over the two periods under examination.

In summary, the evidence presented in Tables 5 and 6 as well as in Figures 1 and 2 indicates that the relative demand for more educated people was a significant determinant of the changes in the returns to education in Argentina during the periods 1996-1999 and 1999-2002. A similar conclusion was reached by Giovagnoli et al. (2004).

Figure 1. Price and quantity changes for 60 groups (1996-1999)

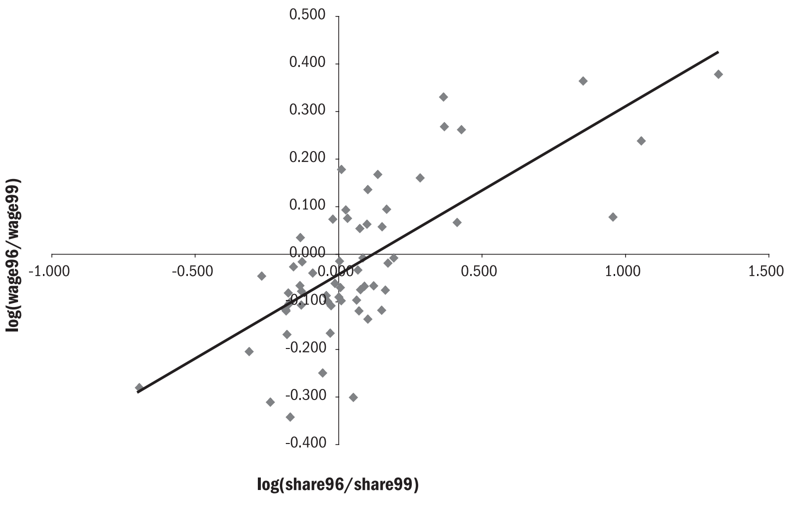
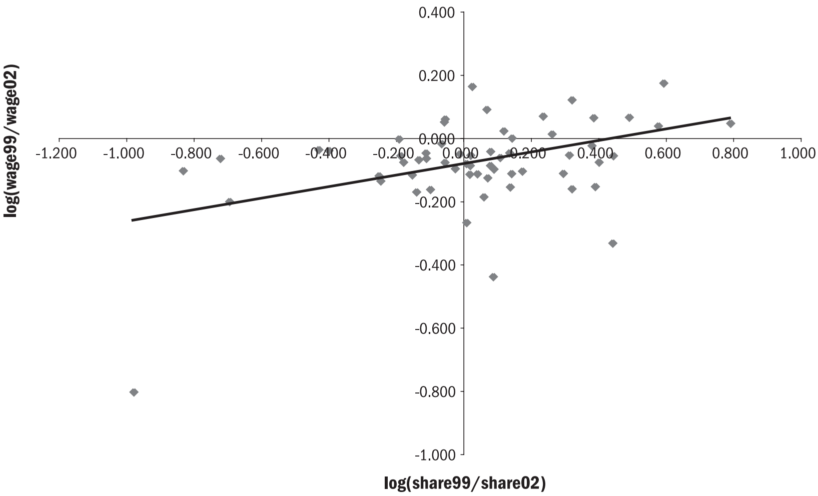


Figure 2. Price and quantity changes for 60 groups (1999-2002)



Finally, one should also consider that the above conclusion may reflect the slow adjustment process of the relative supply of skilled workers in response to a change in the education premium. Although a greater return to skills is expected to induce more people to invest in education, it will take a few years before this pool of more educated individuals enters the labour market. On the other hand, changes in the relative demand are likely to be more rapid as, for instance, they can be related to the expansion of high-technology firms which use more educated workers intensively.

V. Concluding remarks

In this paper we have analysed the changes in the returns to education in Argentina between 1995 and 2003. A standard Mincerain earnings equation was estimated using three different estimation techniques. In addition to the standard OLS method, a Heckman two-stage process was used to account for the sample selection bias arising from possible non-random labour participation. Finally, an IV procedure was also employed in an attempt to control for the potential endogeneity of education.

As in several studies, we find that the IV estimates of the rate of return to education are significantly higher than the corresponding OLS estimates. Additionally, although the estimates associated with these techniques show a different inter-temporal pattern, they concur that returns to education decreased between 1996 and 1999, while they increased in the 1999-2002 period.

Our analysis provides some evidence that the relative demand for more educated people may be an important factor in explaining the observed changes in the returns to education. Between 1996 and 1999 shifts in employment towards industries that intensively employ less-skilled workers seem to be consistent with falling returns to education. By contrast, in the 1999-2002 period shifts in the industrial and occupational composition of employment towards relatively skill intensive sectors are accompanied by a higher education premium.

Appendix

Table A1. Testing differences in the rate of return to education between OLS estimates and OLS estimates corrected for sample selection bias, male and female employees

Year	Statistic
1995	279.93***
1996	352.87***
1997	149.94***
1998	132.30***
1999	131.15***
2000	113.05***
2001	167.03***
2002	196.38***
2003	130.61***

Note: *** Significant at 1 percent level.

The test statistic shown above is computed as $(\alpha_{ols} - \alpha_{ols-corr}) / (\sqrt{s^2_{ols}/n} - \sqrt{s^2_{ols-corr}/n})$, where α_{ols} and $\alpha_{ols-corr}$ refer to the rate of return to education from standard OLS estimates and OLS estimates corrected for sample selection bias, respectively; s_{ols} and $s_{ols-corr}$ refer to the standard errors of the rate of return to education from standard OLS estimates and OLS estimates corrected for sample selection bias respectively and n is the sample size. The test clearly rejects the null hypothesis that the rate of education is the same across the two estimates in all cases.

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